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MANDATORY IFRS REPORTING AND STOCK PRICE INFORMATIVENESS

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Mandatory IFRS Reporting and Stock Price Informativeness

Abstract: In this paper, we examine whether mandatory adoption of IFRS influences the flow of firm-specific information and contributes to stock price informativeness as measured by stock return synchronicity. Using a constant sample of 1,904 mandatory IFRS adopters in 14 EU countries for the period 2003-2007, we find a V-shaped pattern in synchronicity around IFRS adoption, which is consistent with IFRS disclosures revealing new firm-specific information in the adoption period (i.e., a reduction of synchronicity) and subsequently lowering the surprise of future disclosures (i.e., an increase in synchronicity). We also find mandatory IFRS adoption increases analysts' ability to incorporate industry-level information into stock price. However, we are unable to detect a reduction in the private information advantage enjoyed by institutional owners post-IFRS adoption. Moreover, we find the synchronicity effects to be more pronounced for firms in countries with larger differences in local GAAP relative to IFRS. Overall, our evidence yields novel insights on the consequences of mandatory IFRS adoption by investigating its effect on stock price informativeness and the distinctive roles played by a firm's information environment.

Keywords: IFRS, mandatory adoption; stock return synchronicity; information environment.

JEL-codes: F36, G15, M41, M48

Mandatory IFRS Reporting and Stock Price Informativeness

1. Introduction

The mandatory adoption of International Financial Reporting Standards (IFRS) by listed enterprises in the European Union (EU) in 2005 represents a significant regulatory reporting change without historical precedent.[1] In this paper, we provide evidence on how the mandatory adoption of IFRS in the EU influences the incorporation of information into stock prices. Specifically, we examine the behavior of stock return synchronicity (idiosyncratic volatility) around the mandatory adoption of IFRS in the EU. Moreover, we investigate the effect of analysts and institutional investors on stock return synchronicity around mandatory IFRS adoption. We further examine whether differences in accounting standards prior to mandatory IFRS adoption play a significant role in how firm-specific information is incorporated into stock prices around IFRS adoption.

Anecdotal evidence indicates that the first time mandatory adoption of IFRS in the EU in 2005 led to a significant increase in disclosed information because firms not only explained transition effects due to the use of IFRS but also disclosed more footnotes about segments, pensions, share-based payments, and other transactions that were not required to be disclosed under local GAAP (e.g., Hall 2008; Hughes 2008). Reports by Big Four auditing firms on IFRS implementation (e.g., KPMG 2006; Ernst & Young 2007; PwC/IPSOS Mori 2007) highlight increased disclosure and improved comparability as a result of IFRS. We argue that as a result of more information

[1] In the current paper, we use the term IFRS to refer to all standards issued by the International Accounting Standards Committee or the committee's successor, the International Accounting Standards Board, even though the standards issued by the former are typically referred to as International Accounting Standards (IAS).

disclosures required under IFRS and enhanced comparability, the mandatory adoption of IFRS potentially alters the incentives to collect and share information, which affects both common and private information flow and hence stock return synchronicity.

We focus on stock return synchronicity because prior research (e.g., Ferreira and Laux 2007) has identified it as a good candidate for a summary of information flow. More recent research (e.g., Hutton et al. 2009) shows that an earnings-based opacity measure is associated with higher stock return synchronicity (equivalently, lower firm-specific return variation), indicating less revelation of firm-specific information. Moreover, our interest in the pattern of stock return synchronicity around mandatory IFRS adoption stems from prior research (e.g., Tobin 1984; Morck et al. 2000; Durnev et al. 2003; Wurgler 2000) that views stock price as an important signal for asset allocation and regards stock price informativeness as crucial for efficient asset allocation.

Our inquiry is motivated by recent theory which emphasizes the role of information markets on asset price movements. We apply Dasgupta et al.'s (2010) framework to the mandatory IFRS reporting in the EU to derive testable predictions. Dasgupta et al. (2010) present a theoretical model that predicts a decrease in synchronicity at the time more firm-specific new information is disclosed and impounded into stock prices. However, the model also predicts a subsequent increase in synchronicity because the new information allows investors not only to improve their predictions about the occurrence of future firm-specific events but also to incorporate the likelihood of occurrence of these future events into stock prices. Consequently, when these events actually happen in the future, investors react less to such news, making stock prices more synchronous. Consistent with their predictions, Dasgupta et al. (2010)

document this dynamic synchronicity response to two equity-related events, namely seasoned equity offerings and U.S. cross-listings. In our context, mandatory IFRS adoption is the event associated with a significant increase in disclosed information (e.g., KPMG 2006). Similar to Dasgupta et al (2010), we predict a decrease in synchronicity in response to mandatory IFRS adoption, and a subsequent increase in synchronicity in the post-IFRS adoption period.

Next we investigate the influence of two informed market participants, namely financial analysts and institutional investors, on the predicted synchronicity patterns around mandatory IFRS adoption.[2] Prior research (e.g., Piotroski and Roulstone 2004; Ramnath 2002) has found that financial analysts are likely to increase the amount of industry-level information in prices because of their ability to better interpret and disseminate common information across all firms in the industry. Therefore, to the extent that mandatory adoption of IFRS enhanced the comparability of financial reports among firms, we expect the effect of greater analyst activity on stock return synchronicity to be more pronounced in both the year of mandatory IFRS adoption and in the post-IFRS adoption period. In contrast, institutional investors possess an information advantage arising from greater monitoring and increased access to private firm-specific information (Piotroski and Roulstone 2004; Xu and Malkiel 2003). To the extent that mandatory IFRS adoption resulted in more timely and more extensive disclosures, the effect of higher levels of institutional holdings on synchronicity may be attenuated in both the year of mandatory IFRS adoption and in the post-IFRS adoption period because of reduced institutional information advantage.

[2] In Graham et al.'s (2005) survey of 401 financial executives, about 90% of them view institutional investors, followed by analysts, as the most important group in terms of setting company stock price.

Finally, we examine whether the predicted synchronicity patterns depend on the degree to which the accounting rules in a country change with the mandatory switch to IFRS. Firms from countries with a large difference between local GAAP and IFRS are likely to experience a more 'lumpy' information shock at the time of mandatory IFRS adoption than those from countries with a small difference between local GAAP and IFRS. Therefore, the decrease in stock return synchronicity in the year of first-time IFRS reporting and the subsequent increase in synchronicity in the post-IFRS adoption period are likely to be the largest for firms from countries with the large differences in local GAAP versus IFRS.

We use an interrupted time-series design to test our predictions on a constant sample of 1,904 mandatory IFRS adopters with December fiscal year-end across 14 EU countries. Turning to our empirical results, we observe an average 8 percent decrease in stock return synchronicity in the year of mandatory IFRS adoption, and an increase of 36 percent in the post-adoption period (compared to the pre-IFRS period). When we control for synchronicity determinants identified in prior research, we continue to find stock return synchronicity went down in the year of mandatory IFRS adoption and subsequently increased in the post-adoption years to levels higher than in the pre-adoption period. In other words, mandatory adoption of IFRS influenced the flow of firm-specific information and contributed to a V-shaped pattern in synchronicity around the adoption year, a finding that is consistent with the theoretical framework of Dasgupta et al. (2010).

We also find that greater analyst activity had a positive effect on synchronicity but only so in the post-IFRS adoption years, which is consistent with the notion that analysts were able to better interpret and disseminate common information across all firms in the industry after the mandatory adoption of IFRS. However, we find no evidence that higher levels of institutional holdings affected stock return synchronicity differently in the year of mandatory IFRS adoption or in the post-IFRS adoption years, which suggests that the mandatory adoption of IFRS did not alter the private information advantage enjoyed by institutional investors.

Finally, when we split the sample based on the number of differences between local GAAP and IFRS on 21 important accounting rules identified by Bae et al. (2008), we find that the V-shaped synchronicity effect is most pronounced for firms domiciled in countries where local GAAP differs more from IFRS. This result is consistent with mandatory IFRS adoptions contributing the largest effects where information flows are potentially most affected. To further substantiate that our results are due to mandatory IFRS adoption and not to some uncontrolled-for time-varying factor, we partition our sample based on the 2004 IFRS-local GAAP earnings per share reconciliation amounts. Results of this subsample analysis indicate that mandatory IFRS adopters with large earnings per share reconciliations experience the greatest drop in stock return synchronicity during the adoption year, highlighting the importance of cross-sectional differences in earnings numbers resulting from the mandatory IFRS adoption

As an additional robustness check, we test how our measure of stock price informativeness relates to the amount of firm-specific information around earnings announcements. Our results indicate that the larger our synchronicity measures are, the weaker the trading volume and return volatility response around earnings announcements are, suggesting that our synchronicity measures capture the extent of firm-specific information capitalization into stock prices.

Overall, our paper documents the dynamic response of stock return synchronicity to mandatory IFRS adoption and helps to understand one factor that can contribute to the stock price formation process, namely a mandated harmonization of financial reporting standards. Prior empirical research has examined how stock return synchronicity is related to either accounting systems or voluntary IFRS adoptions. However, the evidence on this relation is somewhat mixed. Using a cross-country research design, Morck et al. (2000) find that the sophistication of a country's local GAAP does not explain synchronous stock price movements. Kim and Shi (2010) find that voluntary IFRS adoption in 34 countries over the 1998-2004 time period decreases stock return synchronicity and that this effect is attenuated for firms with high analyst following. To our knowledge, however, no previous research has examined the relation between stock return synchronicity and mandatory IFRS adoption. Our results show that mandatory adoption of IFRS influenced stock return synchronicity in EU countries, which suggests that the first time mandatory adoption of IFRS in 2005 altered capital market information flows.

Our study also contributes to extant research on the consequences of IFRS. Using data from the first few annual reports released under the new regime, Daske et al. (2008) show that market liquidity and equity valuations increase around the time of mandatory introduction of IFRS across 26 countries (including 18 EU countries). However, they find mixed evidence regarding the effects of IFRS on the cost of equity capital. Li (2010) focuses exclusively on EU countries and finds that mandatory adopters experience a significant reduction in the cost of equity capital in the year of mandatory introduction of IFRS. In the current study, we are able to provide evidence on how mandatory IFRS adoption affects information flows over a longer time period.

Finally, our study highlights the role of several key elements of the information environment and how these elements interact with accounting standards. Bushman et al. (2004) conceptualize a firm's information environment as a multifaceted system whose components collectively produce, gather, validate, and disseminate information. In the current study, we document whether two informed market participants, financial analysts and institutional owners, interact with the change in standards to affect stock return synchronicity even further.

The remainder of the paper is organized as follows: Section 2 describes the related literature and develops our testable hypotheses. Section 3 discusses the sample and section 4 details the empirical methods. In section 5, we present empirical results. Section 6 concludes the paper.

2. Hypothesis Development

Grossman and Stiglitz (1980) predict that improving the cost-benefit trade-off on private information collection leads to more extensive informed trading and to more informative pricing. Jin and Myers (2006) develop a theory linking managerial opportunism, transparency, and firm-specific return variation that supports this interpretation. They argue that transparency affects the division of risk bearing between managers and investors. For example, in more transparent firms, insiders take on less firm-specific risk, while outsiders bear less market risk. As a result, more firm-specific disclosure results in a firm's stock price reflecting more firm-specific information and less stock return synchronicity. Jin and Myers (2006) find evidence consistent with this cross-sectional prediction.

In a time series setting, Dasgupta et al. (2010) show that stock return synchronicity can increase when transparency improves. In particular, when the information environment surrounding a firm improves as more firm-specific information is disclosed, stock return synchronicity will initially decrease, which is consistent with Jin and Myers (2006). However, to the extent market participants are able to improve their predictions about the occurrence of future firm-specific events from the disclosure of time-varying firm-specific information (or disclosure of time-invariant information about firm characteristics), stock return synchronicity will subsequently increase.[3] Dasgupta et al. document this dynamic synchronicity pattern for two equity issuance events, namely seasoned equity issues and listing of ADRs, which are associated with significant amounts of new or additional information disclosure (e.g., Lang et al. 2003).

In this study, we examine the mandatory adoption of IFRS in Europe, an event that has been argued to be associated with increased disclosure and enhanced comparability (Daske et al., 2008; Li 2010). EU-listed firms were required to publish their financial reports according to IFRS when they reported their 2005 performance (EC Regulation 1606/2002). Some firms disseminated the information early that year through interim reports, press releases, and documents explaining transition effects, while others waited until the release of the fully IFRS compliant reports (Christensen et al. 2009). Since IFRS typically requires more information to be disclosed, such as footnotes about

[3] While theory predicts the direction of stock return synchronicity over time to the release of new information, it provides little guidance on the magnitude of the effects on synchronicity associated with new information.

segments, pensions and share-based payments, first-time IFRS reporting is analogous to Dasgupta et al.'s characterization of 'lumpy' new information being released.[4] The implication is that the lumpy one-time IFRS information reveals new firm-specific information in the adoption year and enables market participants to improve their predictions about the occurrence of future firm-specific events, thereby lowering the surprise of future disclosures. Using Dasgupta et al.'s framework, the prediction is that stock return synchronicity first decreases in the year of first-time IFRS reporting and subsequently increases. Hypotheses 1a and 1b can be formally stated as follows:

H1a: Stock return synchronicity decreases in the period of first-time IFRS-compliant reporting, ceteris paribus.

H1b: Stock return synchronicity increases in the period following the first-time IFRS-compliant reporting, ceteris paribus.

Next, we investigate the influence of financial analysts on the predicted synchronicity patterns around mandatory IFRS adoption. Prior studies focusing on the information content of analyst recommendations at the firm level provide evidence on the analysts' role in security price formation. For example, empirical studies by Womack (1996) and Jegadeesh et al. (2004) document that analyst recommendations convey useful firm-specific information. Others (e.g., Easley et al. 1998) argue that because analysts make their reports known to a range of investors, the accessibility of these reports may serve to turn private information into public information. Brennan et al. (1993) find that the returns on stocks followed by many analysts lead those of stocks followed by few analysts. More recently, Howe et al. (2009) find that aggregate analyst recommendations

[4] We recognize that mandatory IFRS adoption affects multiple firms in the same time span, potentially influencing market returns, return volatility, and risk factor pricings. In our empirical models, we therefore build in sufficient controls to isolate the price formation effects of lumpy information that enters the market via first-time IFRS reporting.

not only predict future earnings but also have information value for future market and industry returns.

Prior research (e.g., Chan and Hameed 2006; Piotroski and Roulstone 2004; Ramnath 2002) has also found that financial analysts are likely to increase the amount of industry-level information in prices because of their ability to better interpret and disseminate common information across all firms in the industry. In other words, analyst activity – intuitively thought as increasing the firm-specific component of information – actually has a positive impact on stock return synchronicity. Hameed et al. (2010) provide an explanation for this in that widely followed stocks may exhibit more co-movement because they are priced more accurately, and are therefore used to infer values for more opaque stocks.

Further, Horton et al. (2008) and Byard et al. (2010) find that the information environment, as proxied by analyst forecast accuracy and forecast dispersion, improved after the mandatory adoption of IFRS. Thus, to the extent financial analysts become even better in disseminating common information across all firms in the industry after the mandatory IFRS adoption, we expect the effect of greater analyst activity on stock return synchronicity to be more pronounced in both the year of mandatory IFRS adoption and in the post-IFRS adoption period. Hypotheses 2a and 2b can be formally stated as follows:

H2a: The positive stock return synchronicity effect of high analyst activity is intensified in the period of first-time IFRS-compliant reporting, ceteris paribus.

H2b: The positive stock return synchronicity effect of high analyst activity is intensified in the period following the first-time IFRS-compliant reporting, ceteris paribus.

Next, we consider the effect of institutional ownership on stock return synchronicity around IFRS adoption. Institutional trading is another important channel through which information is incorporated into stock prices. Prior research suggests that institutional investors contribute to private information collection and trading because they spend substantial resources on information research (Hartzell and Starks 2003; Chemmanur et al. 2009). Other studies also point to the informational advantage of institutional investors over other investors due to their greater monitoring abilities (Carleton et al. 1998; Nofsinger and Sias 1999). The implication is that institutional investors may reduce stock return synchronicity (equivalently, increase the relative flow of firm-specific information).[5]

In our context, it is possible that a mandatory switch to IFRS may not attenuate private information trading by institutional investors because this investor group is likely to continue to keep its information advantage through active interaction and involvement with companies (Balsam et al. 2002; Collins et al. 2003; Ke and Petroni 2004). In other words, even if mandatory IFRS adoption increases the total information flow released to the market, it does not necessarily eliminate the information advantage of the institutional investors. However, recent research (e.g. Florou and Pope 2009; Yu 2009) finds that institutional investors increased their participations in IFRS adopting firms, suggesting that the information environment after the mandatory adoption of IFRS was perceived as higher quality. To the extent that mandatory IFRS adoption resulted in more timely and more extensive disclosures (Deloitte 2005, KPMG 2006), the effect of higher levels of institutional holdings on synchronicity may be attenuated in both the year of mandatory IFRS adoption and in the post-IFRS adoption period because of reduced institutional information advantage. The net effect of institutional holdings on stock price

[5] This information advantage is different from the one financial analysts experience in that institutional investors act more as firm insiders and may trade on the private information instead of disseminating information to the general public (Piotroski and Roulstone 2004; Xu and Malkiel 2003).

synchronicity around IFRS adoption therefore remains an empirical issue. We formally state hypotheses 3a and 3b as follows:

H3a: The negative stock return synchronicity effect of institutional ownership is attenuated in the period of the first-time IFRS-compliant reporting, ceteris paribus.

H3b: The negative stock return synchronicity effect of institutional ownership is attenuated in the period following the first-time IFRS-compliant reporting, ceteris paribus.

Finally, we consider whether the impact of mandatory IFRS adoption on stock return synchronicity differs with the degree to which the accounting rules in a country change with the mandatory switch to IFRS. We focus on the extent of changes in the accounting rules because the lumpiness of the newly released information under first-time IFRS reporting is likely to be a function of the difference between the old (i.e., local GAAP) and IFRS reporting. Firms from countries with a large difference between local GAAP and IFRS are likely to experience a more 'lumpy' information shock at the time of mandatory IFRS adoption and a reduced post-adoption surprise compared to firms from countries with only a small difference between local GAAP and IFRS. Therefore, the decrease in stock return synchronicity in the year of first-time IFRS reporting and the subsequent increase in stock return synchronicity in the post-IFRS adoption period is likely to be larger for firms from countries where IFRS constitutes a substantial switch from local GAAP. Hypotheses 4a and 4b can be formally stated as follows.

H4a: The decrease in stock return synchronicity in the period of first-time IFRS-compliant reporting is likely to be larger for firms in countries with large local GAAP to IFRS difference than for firms in countries with small local GAAP-IFRS difference, ceteris paribus.

H4b: The increase in stock return synchronicity in the period following the first-time IFRS-compliant reporting is likely to be larger for firms in countries with large local GAAP to IFRS difference than for firms in countries with small local GAAP-IFRS difference, ceteris paribus.

3. Sample

Table 1 summarizes the sample selection procedure. We construct our sample from the list of all firms from 14 EU (EU15 excluding Luxembourg) member countries that are covered by the Worldscope database.[6] For each firm-year observation, we require that there is sufficient data available in Worldscope to compute the financial data items and stock returns used in the empirical tests.

To avoid the confounding effects of changes in firm coverage over time (such as the inclusion in later years of younger, less profitable, and more high-tech firms in the database), we restrict our sample to firms with complete data for each year during the 2003-2007 time period. Because the focus of this study is on the effects of mandatory IFRS adoption, we exclude firms that voluntarily adopted IFRS prior to 2005 or adopted IFRS only after 2007 (such as AIM firms on the London Stock Exchange).[7] We also exclude firms in regulated industries with SIC codes 49 and 62 because stock prices of regulated firms are expected to respond similarly to changes in underlying regulations and economic conditions (Piotroski and Roulstone 2004). Moreover, we exclude EU firms cross-listed in the U.S. because these firms differ from EU firms not cross-listed in the U.S. in the degree of association between accounting data and share prices (Lang et

[6] EU15 refers to the 15 European Union member countries that were members of the EU before the enlargement to 25 (27) countries on May 1, 2004 (January 1, 2007). They include (in alphabetical order): Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden, and United Kingdom.

[7] Moreover, the results for the voluntary IFRS adopters may be confounded if mandatory adoption of IFRS forced these firms to provide more competitive disclosures.

al. 2006).[8] To ensure that the returns for each sample firm are aligned in time for computing our measures of stock return synchronicity, we restrict our sample to firms with December fiscal year-ends. This screening process results in 9,520 firm-year observations (1,904 unique firms) covering 55 two-digit SIC codes.

[Insert Table 1]

The number of firm-year observations covered within a country range from 75 in Austria to 1,815 in the UK. Industries are fairly well represented in the final sample. Largest number of observations (about 20.2 percent) belong to SIC code 3 (Manufacturing). The least number of observations (about 4.8 percent) are from SIC code 8 (Services other than entertainment, food and accommodation).

4. Measurement of Variables and Model Specification

4.1. Measurement of Stock Return Synchronicity

In order to measure stock return synchronicity, we follow the methodology outlined in previous studies (e.g., Morck et al. 2000; Durnev et al., 2003; Piotroski and Roulstone 2004). We consider three alternative specifications of the empirical model using weekly returns for each firm over a 12-month period. First, we regress weekly returns on the current and prior week's value-weighted market return as follows:

$$RET_{i,w} = \alpha + \beta_1 MARET_{i,w} + \beta_2 MARET_{i,w-1} + \varepsilon_{i,w} \qquad (1)$$

where *w* refers to week w, RET is firm-level return compounded weekly, and MARET equals the value-weighted market-wide compounded return that is computed using all firms in the market (except for firm i). Following Piotroski and Roulstone (2004), we

[8] Prior research (e.g., Coffee 1999; Karolyi 2006) has noted that a U.S. listing subjects a cross-listed firm to a more stringent enforcement regime, improves investors' ability to take action through low cost actions such as class actions and derivative actions, and requires it to commit to a higher level of disclosure.

include a lagged market return metric because of the potential for the market information to be incorporated into prices with a delay.

Second, we augment the market model by regressing the firm-level weekly return on the current week's and prior week's value-weighted market and industry-level return ($INDRET_{i,w}$) based on two-digit SIC codes.[9]

$$RET_{i,w} = \alpha + \beta_1 MARET_{i,w} + \beta_2 MARET_{i,w-1} + \beta_3 INDRET_{i,w} + \beta_4 INDRET_{i,w-1} + \varepsilon_{i,w} \qquad (2)$$

where INDRET equals the value-weighted industry-level return for week w using all firms in the industry (excluding firm i), respectively, and where industry is defined based on the same two-digit SIC code as firm i.

Third, we estimate the Fama and French (1996) three-factor model to adjust for exposures to systematic risk arising from a market factor, a size factor, and a value factor. Specifically,

$$RET_{i,w} = \alpha + \beta_1 MARET_{i,w} + \beta_1 SMB_{i\ w} + \beta_1 HML_{i,w} + \varepsilon_{i,w} \qquad (3)$$

where the terms SMB (small minus big) and HML (high minus low) are the weekly returns on zero-investment factor-mimicking portfolios designed to capture size and book-to-market effects, respectively, and all other variables are as defined before. We obtain weekly international SMB and HML factors for EU countries from Bekaert et al. (2009), who compute SMB for each country as the difference between the value-weighted returns of the smallest 30% of firms and the largest 30% of firms within a country. They compute the HML factor in a similar way using high versus low book-to-market stocks.

[9] All results relating to the synchronicity variables based on model (1) and (2) reported in this paper define MARET and INDRET using country-weighted returns. Unreported results using an EU market index and EU industry indices yield inferences similar to those reported in the paper.

We estimate equations (1), (2), and (3) for firm i in year t using weekly returns over the 12-month return window ending 4 months after the fiscal year-end to ensure that earnings-related news flows into the returns during the time-period used for computing our synchronicity measures.[10] In doing so, we require weekly return data to be available for at least 45 weeks in each estimation period. As in other studies, we define stock return synchronicity (SYNCH) as $\log\left(\frac{R^2}{1-R^2}\right)$ where R^2 is the coefficient of determination obtained from one of the three estimation models: market model, industry-augmented market model, and Fama-French 3 factor model. Specifically, we consider three different measures of SYNCH using R^2 from equations (1), (2), and (3), denoted by SYNCH(1), SYNCH(2), and SYNCH(3), respectively. The log transformation changes the R^2 variable, bound by zero and one, into a continuous variable with a more normal distribution. By construction, higher values of this variable reflect higher comovement of firm returns.

4.2. Empirical Models

To examine the relation between stock return synchronicity (i.e., the firm-specific informativeness of stock prices) and the mandatory adoption of IFRS, we estimate the following model:

$$\begin{aligned}SYNCH_{i,t} = {} & \alpha_0 + \beta_1 ADOPT_i + \beta_2 POST\_ADOPT_i + \beta_3 Log(NREV_{i,t}) \\ & + \beta_4 ADOPT_i * log(NREV_{i,t}) + \beta_5 POST\_ADOPT_i * log(NREV_{i,t}) \\ & + \beta_6 INSTIT_{i,t} + \beta_7 ADOPT_i * INSTIT_{i,t} + \beta_8 POST\_ADOPT_i * INSTIT_{i,t}\end{aligned}$$

[10] To assess whether this assumption was reasonable, we checked the annual earnings announcement dates of firms in our sample and found that 95 percent of our firms had an earnings announcement before April 30. To the extent that the release of IFRS-based financial statements could have occurred after April 30, it is possible that some of the IFRS-based financial information may not be reflected in our synchronicity measure, thereby biasing our tests against finding the predicted effects.

$$+ \Phi X + \gamma_j \sum IND_j + \varepsilon_{i,t} \quad (4)$$

Where:

SYNCH	=	a measure of synchronicity of firm-level stock returns.
ADOPT	=	dummy variable equal to 1 if a firm observation relates to the fiscal year ending on December 31, 2005, and 0 otherwise.
POST_ADOPT	=	dummy variable equal to 1 if a firm observation relates to the fiscal year ending on December 31, 2006 (or 2007) and 0 otherwise.
Log(NREV)	=	log of average number of analyst revisions of one-year ahead forecasts of annual earnings during the pre-IFRS adoption period [Source: IBES detailed files].
INSTIT	=	the average proportion of institutional holdings in a firm during the pre-IFRS adoption period [Source: Amadeus].
Φ, X	=	Vector of coefficients and control variables, respectively.
IND	=	industry fixed effects based on one-digit SIC code.
i, j, t	=	firm *i*, industry *j*, and time *t*, respectively.

We utilize a pooled time-series, cross-sectional approach to estimate model (4). Moreover, we use firm clustering to estimate the standard errors of coefficients to account for the correlation of a given firm's residuals over time. That is, *t*-statistics are based on White (1980) robust variance estimates that are adjusted for within-cluster correlation where a firm comprises the cluster (Petersen 2009).[11]

The empirical model (4) includes two main effect variables, ADOPT and POST_ADOPT, to capture the effects of mandatory IFRS adoption relative to the pre-period covering fiscal years ending on December 31, 2003 and December 31, 2004.[12]

[11] This firm clustering approach assumes that the time effect is fixed. In case of the presence of a time effect that is not fixed, the econometrics literature recommends the use of clustering on two dimensions simultaneously (e.g. firm and time). However, the effectiveness of clustering on two dimensions depends on the ability to derive unbiased clustered standard error estimates. For example, Petersen (2009, Figure 5, p. 455) shows that bias in clustered standard error estimates declines with number of clusters, dropping from 27% when there are five years (or clusters) to 3% when there are 40 years. Similarly Gow et al. (2010) show that two-way robust standard errors based on firm and year can produce better inferences when there are at least as few as 10 year clusters. Given that we have 5 years of data and that our empirical model (4) has interactions of analysts and institutions with year effects, there is not enough variation in our year-clusters. As a result, we do not cluster standard errors by time.

[12] To rule out the possibility that the information in IFRS-based financial statements was anticipated prior to the mandatory IFRS adoption, we repeated our empirical tests by deleting the year 2004 in defining the

These two variables are also interacted in model (4) with the two variables representing informed market participants, financial analysts and institutional investors, to assess the role of these participants in affecting stock return synchronicity over time. Empirical model (4) also includes industry dummies to control for potential industry fixed effects along with several firm-specific and country-level control variables, which are discussed in more detail in section 4.3. Appendix 1 provides a detailed description of all variables of interest including the control variables.

Hypothesis H1a is supported if β_1 in model (4) is significantly negative and hypothesis H1b is supported if β_2 in model (4) is significantly positive. In terms of the test of hypotheses H2a and H2b relating to financial analysts, we expect the signs of the coefficients (β_4 and β_5) on the interaction terms, ADOPT*log(NREV) and POST_ADOPT*log(NREV), to be positive. Moreover, for hypotheses H3a and H3b relating to institutional investors, we expect the sign of the coefficients (β_7 and β_8) on the interaction terms, ADOPT*INSTIT and POST_ADOPT*INSTIT also to be positive.

To test hypothesis H4a and H4b, we partition our sample based on the number of differences between the local GAAP of a sample country to IFRS in 2000, and re-estimate empirical model (4) for each partition.[13] We then use a Chow (1960) test to evaluate the differences in the parameter estimates (β_1 and β_2) for the two subsamples.

4.3 Control Variables

We now turn to the main firm- and country-level control variables used in our empirical model (4). These variables, which have been argued or shown in prior research to be related to stock return synchronicity, are as follows.

Pre-IFRS period. Untabulated results using the new benchmark period indicate that our inferences remain unchanged when we use 2003 as the Pre-IFRS period.

[13] We describe the measurement of the local GAAP difference relative to IFRS in section 4.4.

Our first set of control variables are based on a direct transformation of R^2, the variable used to compute our synchronicity measure. Specifically, $R^2 = \beta^2 S_{xx} / (\beta^2 S_{xx} + SSE)$, where β is the stock's co-movement with the market, S_{xx} is the market-wide return variation, and SSE is the idiosyncratic return variation. Thus, we include $\text{Beta}_{(\text{MARKET})}$ as a control when SYNCH(1), which is based on the market model, is the dependent variable. We include both $\text{Beta}_{(\text{MARKET})}$ and $\text{Beta}_{(\text{INDUSTRY})}$ as control variables when SYNCH(2), which is an industry-augmented market model, is used. In case of SYNCH(3), which is based on Fama-French 3-factor model, we include $\text{Beta}_{(\text{MARKET})}$, $\text{Beta}_{(\text{HML})}$, and $\text{Beta}_{(\text{SMB})}$. For all three operationalizations of SYNCH, we introduce log(Total Volatility), which captures the market-wide volatility as an additional variable to control for market uncertainty elicited by mandatory IFRS adoption.

We introduce the natural logarithm of the number of revisions, NREV, to measure the average number of analyst revisions and the variable INSTIT to measure the average number of shares held by institutions (as a fraction of the number of shares outstanding at the beginning of the year). Both these variables are measured in the period prior to mandatory IFRS adoption.[14] Based on prior literature discussed in section 2, we expect a positive relation between NREV and SYNCH and a negative relation between INSTIT and SYNCH.

We include the market value of equity (MCAP) at fiscal year-end to control for a firm's size and its associated effects on stock return synchronicity. To the extent that

[14] We use number of revisions and not analyst following because analyst revision activity is a better indicator of analyst effort compared to the number of unique analysts following a firm (Piotroski and Roulstone 2004). Further, we measure NREV and INSTIT using Pre-IFRS data to avoid the problem of reverse causality because higher analyst activity and higher institutional holdings in 2005 and later may be caused by better disclosures post-IFRS. We thank an anonymous referee for this suggestion.

firm size is positively associated with various dimensions of a firm's information environment not captured by NREV and/or INSTIT, firm size could negatively influence stock return synchronicity. Alternatively, if information acquisition is less costly for large firms, then, in equilibrium, investors may optimally choose to learn more about large firms (Kelly 2007). Thus, firm size could also influence stock return synchronicity positively. Given these conflicting arguments, we do not predict the sign for the variable MCAP in the regressions.

Further, growth opportunities and leverage are likely to affect stock return synchronicity if these characteristics expose a firm to financial distress. We include end of the fiscal year market-to-book ratio (MTB) and the ratio of total debt to total assets (LEVERAGE) as additional control variables. Because of their higher intrinsic risk, firms with more growth opportunities and higher financial leverage are likely to exhibit lower stock return synchronicity. Therefore, we predict a negative relation between SYNCH and both MTB and LEVERAGE.

Synchronicity has been also linked to the age of the firm (Dasgupta et al., 2010). Older firms tend to exhibit more synchronicity than do younger firms because investors are less informed about younger firms. We therefore control for firm age measured as the number of years since its year of listing on the local stock exchange. We expect a positive relation between AGE and SYNCH.

Another determinant of stock return synchronicity is industry concentration (Piotroski and Roulstone 2004). To the extent firms' performances are more interdependent in a concentrated industry, news about an individual firm operating in a concentrated industry is likely to affect its industry peers more. We therefore control for the extent of industry-level concentration by using the Herfindahl index (HERF) for each year and for each industry based on a two-digit SIC classification. HERF is the sum within each industry of the square of each firm's market share based on its revenues relative to the total revenues of the industry in that firm's country of domicile.[15] Higher HERF values imply that market share is concentrated in the hands of a few large firms. Hence, SYNCH is expected to be positively related to HERF. In regression models where SYNCH(2) is the dependent variable, we also include the average number of firms (NIND) used to compute weekly industry-level returns to control for any differences in SYNCH(2) arising from differences in sample size used for estimation purposes (Piotroski and Roulstone, 2004).

Finally, model (4) also includes two country-level characteristics that are likely to affect stock return synchronicity. The first country-level variable we control for is the variation in economic conditions across countries (i.e., the state of the economy) by including the annual real GDP growth rate (GDPG). Data on GDPG is obtained from World Bank (2008). Kearney and Poti (2008) report intertemporal variation in idiosyncratic volatility in EU markets, suggesting that synchronicity may (at least partially) reflect business cycle patterns. We do not predict a particular direction for the variable GDPG in the regressions. We also include a country-specific measure for the ease with which companies can raise capital (ACCESS). This variable, obtained from World Bank (2006), is available only for the year 2004. It ranges between 0 (very hard)

[15] Results remain qualitatively unchanged, however, when we define HERF at the EU-year level. Separately, we recognize that the change in accounting standards worsens the problem of obtaining comparable data over time, potentially affecting the choice and computation of control variables based on financial statement data. As a sensitivity test, we recomputed HERF using number of employees. The correlation between revenues-based HERF and HERF based on the number of employees is 0.92 ($p < 0.01$).

and 7 (very easy).[16] Morck et al. (2000) note that better access to capital stimulates informed trading attributable to firm-specific price changes, resulting in less stock return synchronicity. Thus, we expect a negative relation between ACCESS and our SYNCH measures.

4.4 Measurement of Local GAAP Differences Relative to IFRS

We use the country-level data reported in Bae et al. (2008, Table 1) to derive a measure of differences in local GAAP and IFRS for each country in our sample. Bae et al. (2008) rely on a comprehensive survey (Nobes 2001) to identify differences in 21 key accounting items for each country in their analyses.[17] A country is deemed to have local GAAP similar to IFRS for an item listed in Table 1 of Bae et al. (2008) if it conforms to IFRS for that item. We then assign the country a 'GAAP difference' score of zero for that item. In contrast, if a country is deemed to have local GAAP different than IFRS for an item listed in Table 1 of Bae et al. (2008), we assign the country a 'GAAP difference' score of one for that item. This procedure is repeated for all 21 accounting items on Bae et al.'s (2008) list and the total 'GAAP difference' score is the sum of the scores for that specific country across all 21 items. This score is our primary measure of local accounting standard differences and has a theoretical range from zero to 21. We denote

[16] We also estimated regression models by substituting ACCESS by an index covering a variety of size aspects (market capitalization to GDP; value traded to GDP; and turnover ratio) of a country's equity market (World Bank 2008), and by property rights index derived by Morck et al. (2000). Results of these alternative estimations did not alter our inferences.

[17] Bae et al. (2008) use strict criteria for identifying GAAP differences and select only 21 items out of a longer list of 80 original key accounting issues. Examples of such items relate to the recognition and measurement of financial instruments (IAS 32/39); impairment losses (IAS 22/38); provisions (IAS 37); employee benefit liabilities (IAS 19); capitalization of research and development expenses and internally generated intangible assets (IAS 38); disclosure of related party transactions (IAS 24); presentation of a statement of changes in equity (IAS 1) and a statement of cash flows (IAS 7). Our results remain qualitatively unchanged if we use the more extensive list of accounting GAAP differences as reported in Ding et al. (2007).

this measure of GAAP differences as DIFF_GAAP. By construction, higher values of this variable reflect greater difference in a country's local GAAP relative to IFRS. In our sample of countries, the DIFF_GAAP variable ranges between a minimum of 1 for Ireland and the UK to a maximum of 17 out of 21 for Greece. For testing purposes, we classify a country as having a large local GAAP-IFRS difference if its DIFF_GAAP value is more than the sample median value of 12.[18]

5. Empirical Results

5.1 Descriptive Statistics

Panel A of Figure 1 shows Box-Whisker plots of the synchronicity measure SYNCH(2), based on the industry-augmented market model, to highlight the evolution in synchronicity for the whole EU market over three distinct time periods.[19] Similarly, panel B of Figure 1 shows the evolution in SYNCH(2) across two subsamples of countries classified by local GAAP-IFRS differences. In these figures, period 1 refers to the years 2003 and 2004 when IFRS adoption was not mandatory. Period 2 covers the year 2005 when mandatory IFRS adoption was effective and firms produced their first IFRS-compliant information, and period 3 refers to the years following first IFRS-compliant financial statements (2006 and 2007).

[Insert Figure 1 and Table 2]

[18] The 8 countries classified in the large local GAAP-IFRS difference category are Austria, Belgium, Finland, France, Greece, Italy, Portugal, and Spain, and the 6 countries classified in the small local GAAP-IFRS difference category are Denmark, Germany, Ireland, Netherlands, Sweden, and UK.

[19] We choose to report only SYNCH(2) box-whisker plots for the reason of brevity. Unreported results using the other two SYNCH measures, SYNCH(1) and SYNCH(3), are very similar. However, it is noteworthy that SYNCH(2) measure, based on the industry-augmented market model, on average, explains the highest percentage (25.7 percent) of individual stock returns. SYNCH(1), based on the market model, explains the least (8.2 percent), and SYNCH(3) based on Fama-French 3 factor model explains 14.6 percent. Further, pairwise correlations between all SYNCH measures are fairly high and vary between a minimum of 0.603 [SYNCH(1) and (3)] and a maximum of 0.806 [SYNCH(1) and (2)].

For the whole EU market, SYNCH(2) exhibits first a decreasing pattern in 2005 (a 8% decrease from -1.79 in 2003-04 to -1.93 in 2005) and then an increasing pattern in the 2006-07 time period (a 36% increase from -1.79 in 2003-04 to -1.15 in 2006 and 2007). A similar V-shaped pattern in SYNCH(2) emerges for the subsample of firms from countries with large differences in local GAAP to IFRS (See Panel B of Figure 1). Test statistics based on Mann-Whitney test for difference in median values of SYNCH(2) around the mandatory IFRS adoption are always statistically significant ($p<0.01$) for the whole EU market and for the subsample of firms from countries with large differences in local GAAP to IFRS. On a univariate basis, the results for the whole EU market provide initial evidence consistent with our hypotheses H1a and H1b, which predict synchronicity to first decrease in the year when mandatory IFRS adoption became effective, and then to increase after the mandatory IFRS adoption.

Moreover, panel B of Figure 1 shows that the SYNCH(2) variable for the subsample of firms from countries with small local GAAP-IFRS difference does not exhibit a decreasing pattern in the year of mandatory IFRS adoption, even though the median SYNCH(2) value in period 3 (the years after the first year of mandatory IFRS adoption) is the highest. To examine whether a particular country is driving the subsample results, Table 2 reports the evolutions in SYNCH(2) over time for each country grouped into either the large or the small local GAAP-IFRS difference subsample. Each of the eight EU countries included in the large local GAAP-IFRS category exhibit a V-shaped pattern in the evolution of SYNCH. However, this V-shaped pattern prevails for 4 of the 6 EU countries classified into the subsample of countries with small local GAAP-IFRS difference. The exceptions are Denmark and Ireland. Taken together, the evidence for subsamples classified by the local GAAP-IFRS difference is consistent with the underlying premise of hypotheses H4a and H4b that the lumpy information shock is most pronounced in countries where IFRS may impact financial reports the most.

Table 3 presents descriptive statistics for selected variables used in our empirical tests. Median SYNCH values are highest for SYNCH(2), suggesting that the industry-augmented market model (2) explains most of the variation in firm-level returns. Further, the mean values of each of the three SYNCH measures are slightly lower than the median values, indicating that the distributions of these variables are a little left-skewed. The mean number of analyst earnings forecast revisions is 11.74; for 25 percent of the firms analysts do more than 12 revisions per year. The mean percentage of institutional holdings in a firm is 6.53 percent. Mean firm size (as measured by the market value of equity or MCAP) is 1.254 billion Euros.[20] Sample firms have a mean (median) market-to-book ratio of 2.49 (1.73). Financial leverage measured as long term debt to total assets (LEVERAGE) is below 72 percent for 95 percent of the observations. The median firm is listed for 14 years when it enters our analysis. There is considerable cross-sectional variation in the HERF value and in the mean number of firms used in calculating the weekly industry returns. The mean real annual GDP growth rate over the period of our study (2003-2007) is 2.49 percent. The average ACCESS index value is 4.98 (on a scale of 0 for very hard to 7 for very easy), and only is below 4 for about 5

[20] This mean value is smaller than that reported by Li (2010) for her sample, which is not surprising because the requirement to compute ex ante cost of capital biased her sample towards larger firms with substantial analyst following.

percent of observations; all of which indicates that our sample of EU countries exhibit easier access to capital than average.

[Insert Table 3]

5.2 Correlations

Table 4 presents the Pearson/Spearman pairwise correlations among variables used in model (4).[21] The correlation between SYNCH(2) and ADOPT is negative and statistically significant ($p<0.01$), suggesting that that there is less stock return synchronicity (i.e., there is more firm-specific information in stock prices) during the first year when IFRS were mandated. In contrast, SYNCH(2) and POST_ADOPT are positively correlated, suggesting that there is more stock return synchronicity (i.e., there is less firm-specific information in stock prices) during the two years following the first IFRS reports. These correlations are consistent with the evolutions in the synchronicity measure documented in Figure 1. Moreover, SYNCH(2) has statistically significant and positive correlations exceeding 0.20 with $Beta_{(MARKET)}$, $Beta_{(INDUSTRY)}$, log(Total Volatility), log(NREV), MCAP, and GDPG.

[Insert Table 4]

The highest (absolute) pairwise Pearson/Spearman correlation among the control variables is 0.61/0.64 (between variables NREV and MCAP), indicating that especially larger firms experience more analyst revision activity. Moreover, the correlations of MCAP with both AGE and INSTIT are positive, suggesting that larger firms are older and have more institutional holdings. To examine whether these correlations are problematic in regression estimations, we diagnose multicollinearity in the regressions

[21] For the reason of brevity, we only report correlations of variables with SYNCH(2). We note that untabulated correlations of variables in model (4) with SYNCH(1) and SYNCH(3) yield similar patterns.

using variance inflation factors (VIFs). Overall, these VIFs are low, suggesting that collinearity is unlikely to be a significant issue in interpreting the regression results.

5.3 Regression Results: Effect on Synchronicity

Table 5 reports regression results using stock return synchronicity as the dependent variable. The dependent variable in columns labeled [1] – [3] is SYNCH(1), which is based on the market model. The dependent variable in columns [4] and [5] are SYNCH(2) and SYNCH(3) based on industry-augmented market model and Fama-French 3-factor model, respectively. The first two columns use alternative specifications of model (4). In particular, the specification in column [1] ignores the role of analysts and institutions and focuses on the main effects of our test variables ADOPT and POST_ADOPT. The specification in column [2] introduces the main effects of analysts and institutions activities as additional control variables. The remaining columns report results for model (4) that considers both the role of financial analysts and institutions around the mandatory IFRS adoption. We do not report the industry fixed effects for brevity. As reported in Table 4, the explanatory power of the models ranges between 0.365 and 0.522.

[Insert Table 5]

In Table 5, the coefficient on ADOPT is negative and statistically significant ($p<0.01$) in all columns, indicating that stock return synchronicity declined with the mandatory adoption of IFRS. In terms of economic magnitude, the mandatory switch to IFRS reduced stock return synchronicity by approximately 5.7 percent (the coefficient of -0.116 on the ADOPT variable divided by the coefficient of -2.03 on the intercept reported in the first column). Since the firm-specific informativeness of stock prices varies inversely with stock return synchronicity, our findings suggest that the firm-specific information in stock prices increased in the first year of the mandatory adoption of IFRS. In contrast, the coefficient on POST_ADOPT is positive and statistically significant ($p<0.01$), indicating that stock return synchronicity increased in the years after the mandatory adoption of IFRS. This positive coefficient represents an increase in stock return synchronicity in the post-IFRS adoption years of about 9.3 percent (compared to the pre IFRS period; computed by dividing the coefficient of 0.19 on the POST_ADOPT variable by the coefficient of -2.03 on the intercept in the first column). Collectively, these results are consistent with the argument that the lumpy information component originating from mandatory IFRS adoption resulted in stock prices behaving more idiosyncratically when the new information was revealed to the market for the first time and subsequently lowered the surprise component in future earnings shocks in the POST_ADOPT period.

We corroborate the main effects of analyst activity and institutional holdings on SYNCH for our EU sample in column [2] and test for hypotheses H2a-H3b in columns labeled [3], [4] and [5]. The coefficients on variable log(NREV) are positive and statistically significant ($p<0.01$), suggesting that analyst activities exert a positive impact on stock return synchronicity. This result is consistent with Piotroski and Roulstone (2004) who find that analysts in the U.S. increase the amount of industry-level information in prices. The interaction of ADOPT with log(NREV) is not statistically significant ($p>0.10$), even though the unreported F-statistic indicates that the sum of the coefficients of Log(NREV) and its interaction with ADOPT is positive and statistically significant ($p<0.01$). The absence of an incremental effect for financial analysts in the year of mandatory IFRS adoption is not supportive of our hypothesis H2a. However, the interaction between POST_ADOPT and log(NREV) is statistically significant ($p<0.01$) with a positive sign, which is consistent with our hypothesis H2b. These results suggest that analysts contributed to more stock return synchronicity only after the first year of mandatory IFRS adoptions. Overall, this result is similar to the evidence in Kim and Shi (2010), who report that the synchronicity-reducing effect of voluntary IFRS adoption is attenuated for firms with high analyst following.

As predicted, the coefficients on the variable, INSTIT, are negative and statistically significant ($p<0.01$), which suggests that more institutional ownership lowers stock return synchronicity. This result is also consistent with Xu and Malkiel (2003) and Piotroski and Roulstone (2004) finding that institutions possess an information advantage and that they are able to increase the relative flow of firm-specific information through their activities. However, the interaction of INSTIT with both ADOPT and POST_ADOPT are not statistically significant ($p>0.10$), indicating that the negative relation between synchronicity and institutional ownership which existed in the pre-IFRS period is not affected in later periods. These results suggest that the mandatory adoption of IFRS did not alter private information collection by institutional investors. This evidence fails to support our predictions under hypothesis H3a and H3b.

Table 5 also shows that there are several other significant determinants of synchronicity for EU firms. The coefficients on risk proxies, $Beta_{(MARKET)}$, and log(Total Volatility) are significantly positive across all columns. This suggests that higher systematic risk and higher market volatility lowers the incorporation of firm-specific information into stock prices, and thus increases stock price synchronicity. Other risk factors such as $\text{Beta}_{(\text{INDUSTRY})}$, $\text{Beta}_{(\text{HML})}$, $\text{Beta}_{(\text{SMB})}$, are also significantly related to our synchronicity measures. Consistent with the U.S. finding of Piotroski and Roulstone (2004), the coefficients on MCAP are significantly positive across all columns. This indicates that in EU countries, stock prices of large firms tend to move together with the market to a greater extent than do stock prices of small firms. Moreover, the coefficients on MTB are significantly negative across all columns, suggesting that firms with high growth opportunities tend to have more firm-specific information incorporated into their share prices, and thus exhibit a lower level of stock price synchronicity. Somewhat surprisingly, the coefficients for LEVERAGE, Log(AGE) and HERF are insignificant across all columns. Finally, stock return synchronicity increases with higher GDP growth rates and it decreases with the variable ACCESS measuring the ease with which firms domiciled in a country can raise capital.

5.4 Regression Results: Stock Return Synchronicity and Earnings Informativeness

While important, our evidence on lower stock return synchronicity around IFRS adoption could alternatively be interpreted as reflecting more noise in firm-level stock returns as opposed to higher stock price informativeness. In other words, the portion of stock returns unexplained by market-wide, industry-level and risk-adjusted returns could still be merely measuring statistical noise. To disentangle the noise versus information component story in our main findings, we perform an additional test to examine if our measure of stock return synchronicity captures the amount of firm-specific information (relative to market-wide and/or industry-wide information) impounded into observed stock prices in the EU market by focusing on a short-window around earnings announcements, which are viewed as important events for value-relevant firm-specific information. If SYNCH correctly captures the extent of firm-specific information capitalization into stock prices, we expect the trading volume and return volatility response to annual earnings announcements to be weaker for firms with high SYNCH than for firms with low SYNCH.

To provide empirical evidence on this issue, we first estimate abnormal trading volume and abnormal return volatility around earnings announcements for a subsample of firms with annual earnings announcement dates available on Worldscope database and then regress these metrics on our synchronicity measures and other factors likely to be associated with abnormal volume and volatility. Additional data restrictions yield a constant sample of 8,250 (1,650 unique firms) and 7,940 (1,588 unique firms) firm-year observations for the abnormal volume and volatility regressions, respectively.

Following Landsman and Maydew (2002), we compute cumulative abnormal volume ($CAVOL_i$) for each firm i around the annual earnings announcement date as the sum of abnormal volume ($AVOL_{id}$) defined as (V_{id} - $\overline{V_i}$)/σ_i, where V_{id} is equal to the number of shares of firm i traded during day d (d = -1, 0, +1) relative to earnings announcement day (d=0), divided by shares outstanding of firm i during day d; and $\overline{V_i}$ and σ_i are the mean and standard deviation, respectively, of daily trading volume for firm i in the period d-345 to d-20 and d+20 to d+345.[22] Analogously, the cumulative abnormal volatility measure ($CAVAR_i$) for each firm around the annual earnings

[22] In estimating the values of $\overline{V_i}$ and σ_i,, we exclude the 20 days before and after each of the other earnings announcements contained within this estimation window. There is some variation with regard to reporting frequency both across and within countries: some firms report on an annual, semi-annual, or quarterly basis. To control for these differences, we exclude each quarterly, semi-annual or annual earnings announcement in the estimation period surrounding each annual earnings announcement, and require at least 100 trading days to compute $\overline{V_i}$ and σ_i.

announcement date is the sum of abnormal volatility ($AVAR_{id}$) defined as u_{id}/σ_i^2, where u_{id} is daily market model-adjusted return computed as $R_{id} - (\alpha_i + \beta_i R_{md})$, R_{id} is the raw return of firm i on day d, R_{md} is equally-weighted return of market for day d, α_i and β_i are firm i's market model parameter estimates, and σ_i^2 is the variance of firm i's market model adjusted returns, each of which is calculated during the period d-345 to d-20 and d+20 to d+345. Unlike the $AVOL_{id}$ measure, $AVAR_{id}$ is always positive, where values between zero and one are indicative of smaller than normal volatility and values greater than one reflect higher abnormal volatility (Landsman and Maydew 2002).

To test the relation between earnings informativeness and synchronicity, we employ the following model: [23]

$$
\begin{aligned}
CAVOL_i \text{ (or } CAVAR_i) = \; & \lambda_0 + \lambda_1 SYNCH_i(2) + \lambda_2 MCAP_i + \lambda_3 MTB_i \\
& + \lambda_4 D\_NEGEPS_i + \lambda_5 LEVERAGE_i + \lambda_4 Beta_{(MARKET)} \\
& + \lambda_7 Beta_{(INDUSTRY)} + \lambda_8 LOG(NREV)_i + \lambda_9 INSTIT_i \\
& + \lambda_j \sum IND_j + \lambda_c \sum CTRY_c + \varepsilon_i \qquad (5)
\end{aligned}
$$

where D_NEGEPS is equal to one for a firm reporting a loss, and 0 otherwise; CTRY refers to country fixed effects, and all other variables are as defined before. In all estimations, we compute *t*-statistics based on White (1980) robust variance estimates that are adjusted for within-cluster correlation where a firm comprises the cluster (Petersen 2009). To alleviate a concern over the possibility that the regression results for model (5) are unduly influenced by a small number of outlier values of SYNCH(2) and/or SYNCH(2) values are measured with error, we also use DR_SYNCH(2) instead of SYNCH(2) in alternative regression specifications. To obtain DR_SYNCH(2), we

[23] For brevity, we tabulate results using SYNCH(2) as the measure of synchronicity. Unreported results using SYNCH(1) or SYNCH(3) yielded similar inferences.

classify our firms into deciles based upon the ranked values of SYNCH(2) in each sample year, with zero representing the smallest decile of SYNCH(2) and nine representing the largest. We then scale the decile ranks to range between zero and one (Landsman and Maydew 2002).

Table 6 presents the regression results using cumulative abnormal trading volume (CAVOL) in columns [1] and [2] and cumulative abnormal return volatility (CAVAR) in columns [3] and [4] as the dependent variables. The coefficients on our synchronicity measures, SYNCH(2) and DR_SYNCH(2), are negative and statistically significant ($p < 0.05$) in all four columns, after controlling for firm size, market-to-book, presence of a loss, financial leverage, systematic risk factors, number of analyst revisions, and fraction of shares held by institutions. These negative relations confirm that high values of SYNCH(2) for our sample firms have low earnings informativeness in terms of both abnormal trading volume and abnormal return volatility around annual earnings announcements. The above findings corroborate the view that our measure of stock price synchronicity is able to capture the extent of firm-specific information capitalization into stock prices in the EU market.

[Insert Table 6]

5.5 Subsample Regression Results Based on local GAAP-IFRS Differences

Next, we re-estimate empirical model (4) for subsets of data to determine whether the shifts in synchronicity were similar for firms coming from countries with large versus small differences in local GAAP relative to IFRS. The regression results using SYNCH(2) as the dependent variable are presented for the two subsamples in Table 7.[24]

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[24] Unreported results using SYNCH (1) and SYNCH(3) as the dependent variable for the two partitions yielded inferences similar to those reported in Table 7.

[Insert Table 7]

The coefficient on the variable ADOPT is statistically significant with a negative sign for the large DIFF_GAAP subsample only, indicating that the synchronicity drop is especially caused by the lumpy information shocks in countries where local GAAP differed substantially from IFRS. Furthermore, the F-statistic for testing the equality of the coefficients on ADOPT for the two subsamples is statistically significant ($p<0.01$), which is consistent with hypothesis H4a. These results suggest that the drop in synchronicity after the mandatory IFRS adoption is higher for firms in countries with large local GAAP-IFRS difference than that for firms in countries with small local GAAP-IFRS difference.

Furthermore, the variable POST_ADOPT is significant with a positive sign for both subsamples, although the coefficient is of a lower statistical significance ($p<0.10$) for firms in countries with small differences in local GAAP to IFRS. The F-statistic for testing the equality of the coefficients on POST_ADOPT for the two subsamples is statistically significant ($p<0.01$), suggesting that the increase in synchronicity in the POST_ADOPT period is especially observed in countries with large differences in local GAAP rather than for firms in countries with small differences in local GAAP.

Taken together, these results suggest that the lumpy IFRS information shock and the subsequent increase in synchronicity were merely caused by firms from countries where large differences existed in local GAAP versus IFRS. This evidence provides supports for both H4a and H4b.

5.6 Subsample Regression Results Classified by the Magnitude of IFRS Restatements

To further substantiate that the observed synchronicity decline in the year of mandatory IFRS adoption reported in Table 5 is indeed attributable to the switch to IFRS, we examine how cross-sectional variation in firm-specific local GAAP to IFRS earnings reconciliations affect the IFRS adoption-synchronicity relation in the year of mandatory IFRS adoption. All EU firms were required to disclose reconciliations between net income under local GAAP and net income under IFRS for their 2004 accounts in the first mandatory IFRS annual reports.[25] We first derive an *ex post* measure of the impact of the IFRS adoption by calculating the IFRS to local GAAP difference as the absolute value of 2004 local GAAP earnings per share (EPS) minus the reconciled 2004 IFRS EPS, scaled by local EPS (i.e., $|(EPS_{LOCAL04} - EPS_{IFRS04})/ EPS_{LOCAL04}|$.[26] Based on this ex post measure, we sort our sample firms into 3 equal-sized portfolios. The mean values of the restatement difference for the three portfolios are 0.028, 0.179, and 2.260, respectively. Given that a large relative restatement difference reflects a large lumpy information effect in a firm's first IFRS reporting, we expect the synchronicity effect in the year of mandatory IFRS adoption to be more pronounced for a firm that experiences a larger restatement effect.

To test this conjecture, we modify model (4) by focusing only on the Pre-IFRS period and the year of mandatory IFRS adoption for the three portfolios and report the separate regression results using SYNCH(2) as the dependent variable for each of the

[25] Although EU firms were also required to disclose a reconciliation of book value of equity and cash flows, we focus on the reconciliation of earnings because such reconciliations are most often included in firms' first-time disclosure of quantitative information regarding the impact of IFRS. Christensen et al. (2009) show significant stock market reactions for a sample of UK firms that release earnings reconciliations early. In the same vein, Horton and Serafeim (2008) find significant negative abnormal returns and positive trading activity for firms reporting a negative reconciliation adjustment on UK GAAP earnings.

[26] We were able to compute this ex post measure for 1,517 firms. Results (available upon request) are robust to various alternative scalars such as total assets, sales and beginning-of-the-year market value.

three groups in Table 8. Note that there are 505 firms in the portfolios labeled as small and large, and that there are 507 firms (1,521 firm-year observations) in the medium portfolio. The ADOPT variable is negative and statistically significant ($p < 0.10$) only for the sub-sample of firms with the largest differences in local GAAP versus IFRS restated EPS. Unreported F-statistics indicate that the absolute value of the magnitude of the coefficients on ADOPT are the largest for the firms in the LARGE restatement group. Again, this evidence is consistent with the conjecture that stock price informativeness undergoes the largest effects when the lumpy information associated with mandatory IFRS adoption is highest. Overall, these results show that stock return synchronicity effects of mandatory IFRS adoptions in the year of adoption vary cross-sectionally and are related to a specific proxy capturing the size of the lumpy information component, namely the difference in the 2004 local GAAP to IFRS earnings reconciliations.

[Insert Table 8]

6. Conclusions

The recent move towards the introduction of a single set of accounting standards such as IFRS is a regulatory reporting change without historical precedent. In this paper, we examine the extent to which mandatory IFRS adoption affects the information flow and contributes to stock price informativeness in the year 2005 when IFRS became mandatory in EU and in the post-IFRS adoption years (2006-2007) relative to pre-IFRS years (2003-2004).

Using a carefully constructed sample of 1,904 EU firms over the period from 2003 to 2007, we find that the stock return synchronicity decreased in the year of mandatory IFRS adoption, but subsequently increased in the post-adoption years to levels higher than the pre-adoption period. This result is consistent with the theoretical prediction that mandatory adoption of IFRS at first is likely to increase the private information flow entering into the stock price formation process and to reduce subsequently the surprise effects of future information releases. Moreover, we find that analysts' activity led to more stock return synchronicity after the mandatory switch to IFRS, which is consistent with the notion that IFRS helped financial analysts in interpreting and disseminating common information across all firms in the industry. However, we find no evidence that higher levels of institutional holdings affected stock return synchronicity differently in the year of mandatory IFRS adoption or in the post-IFRS adoption years, suggesting that the mandatory adoption of IFRS did not alter the private information advantage enjoyed by institutional investors. We also find that the V-shaped synchronicity effect is most pronounced for firms domiciled in countries where local GAAP differs more from IFRS, which is consistent with mandatory IFRS adoptions contributing the largest effects where information flows are potentially most affected. Taken together, these results yield novel insights into how a mandated financial reporting harmonization process shapes capital market information flows including the distinctive roles played by a firm's information environment on this particular process. Overall, our paper contributes to the ongoing debate on the economic benefits and costs of a mandatory accounting change.

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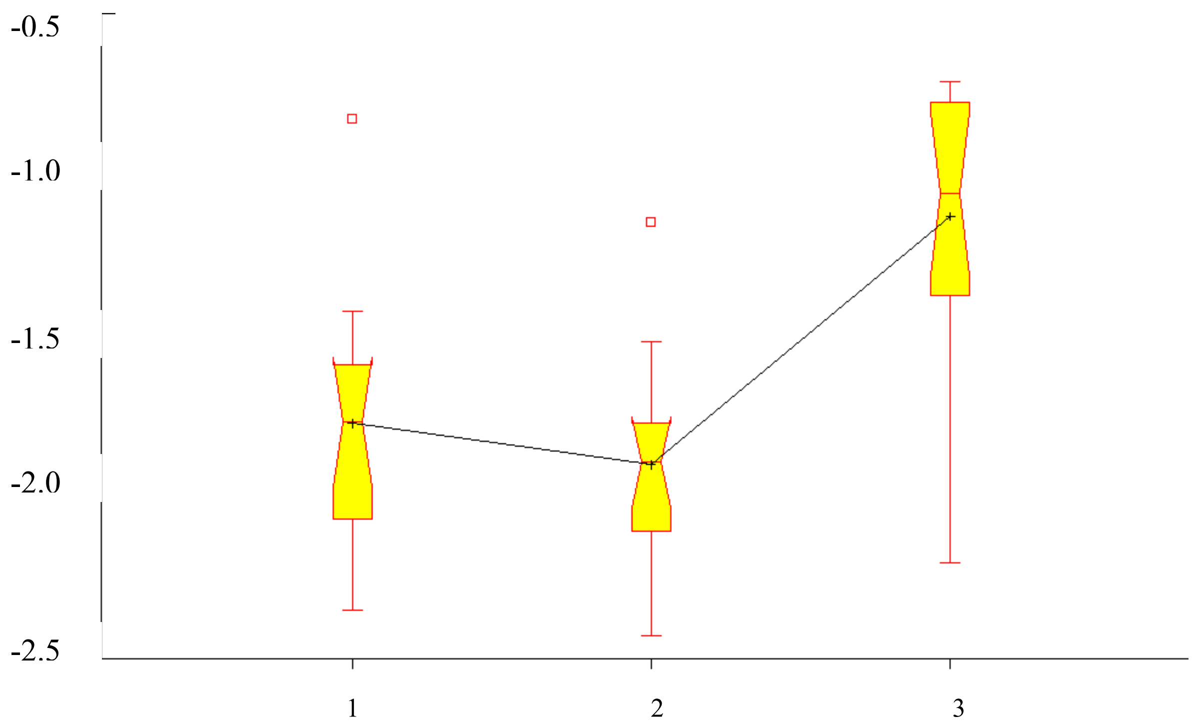
Figure 1: Median Synchronicity Evolution

This Figure reports Box-Whisker plots of country-median value of SYNCH(2) across three periods and tests of statistical differences between country-median synchronicity measure (z-value; Mann-Whitney test) over the three periods. Panel A contains values for all countries (pooled) and Panel B is for countries with above (left-hand side) and below (right-hand side) sample median country-level differences between local GAAP and IFRS. DIFF_GAAP is LARGE (SMALL) if local GAAP differs substantially from IFRS on more than (less than) 12 [i.e., the sample median] of 21 key accounting items as in Bae et al. (2008). Period 1 refers to 2003 and 2004 when IFRS adoption was not mandatory. Period 2 refers to 2005 when IFRS adoption became mandatory. Period 3 covers years after the first year of mandatory IFRS adoption (2006 and 2007). Significance levels are reported in brackets. The Variable SYNCH(2) is defined in Appendix 1.

Panel A: All Countries

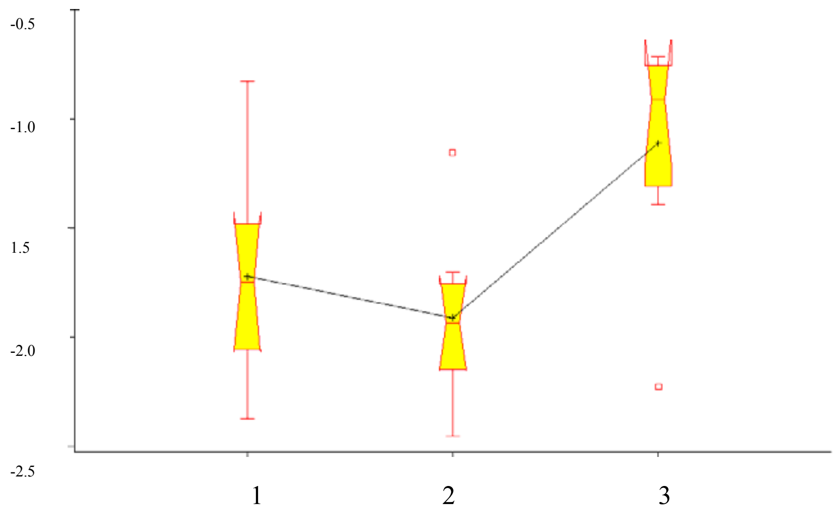


Median	–1.79	–1.93	–1.15

	Period 1 vs. Period 2	**Period 2 vs. Period 3**	**Period 1 vs. Period 3**
z-value	3.94 (<0.01)	23.58 (<0.01)	22.92 (<0.01)

Panel B: Country Classified by Differences in Local GAAP to IFRS

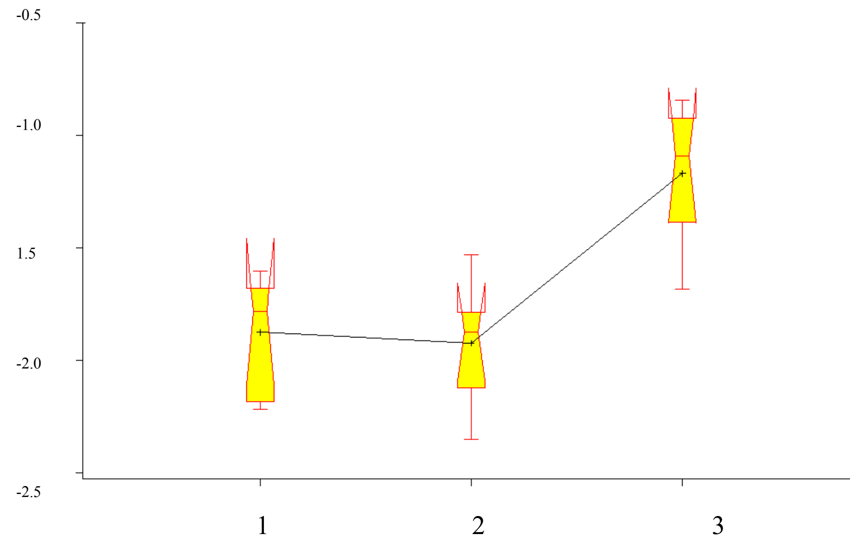
Countries with *DIFF_GAAP = LARGE*



Median –1.64 –1.85 –1.01

	Period 1 vs. Period 2	Period 2 vs. Period 3	Period 1 vs. Period 3
z-value	4.59 (<0.01)	18.12 (<0.01)	16.71 (<0.01)

Countries with *DIFF_GAAP = SMALL*



Median –1.91 –1.98 –1.29

Period 1 vs. Period 2	Period 2 vs. Period 3	Period 1 vs. Period 3
1.08 (0.27)	14.47 (<0.01)	16.79 (<0.01)

Table 1: Sample Selection

Criterion	Number of Firm-years	Number of Distinct Firms
EU15 Country firms, with financial statement items and stock market information available in the period 2003-2007:	20,552	5,234
(1) Minus firms with less than 5 years of relevant information available:	-5,172	-2,068
=	15,380	3,166
(2) Minus voluntary IFRS adopters (pre-2005):	-1,225	-335
=	14,155	2,831
(3) Minus post 2005 IFRS adopters:	-2,860	-572
=	11,295	2,259
(4) Minus regulated industries (SIC49, SIC62):	-430	-86
=	10,865	2,173
(5) Minus US ADR firms:	-510	-102
=	10,355	2,071
(5) Minus non-December 31 FYE firms	-835	-167
=	9,520	1,904

Table 2: Descriptive Statistics of SYNCH(2) by Country-Period

DIFF_GAAP = LARGE						DIFF_GAAP = SMALL					
Country	**Period**	**N**	**Median**	**Mean**	**Std.**	**Country**	**Period**	**N**	**Median**	**Mean**	**Std.**
AUSTRIA	1	30	-2.377	-2.252	0.785	DENMARK	1	170	-2.186	-2.328	1.176
	2	15	-2.454	-2.393	1.017		2	85	-1.786	-1.927	1.160
	3	30	-2.228	-2.042	1.146		3	170	-1.098	-1.265	1.298
BELGIUM	1	118	-2.089	-2.256	1.375	GERMANY	1	376	-2.217	-2.311	1.081
	2	59	-2.126	-2.345	1.248		2	188	-2.353	-2.418	1.027
	3	118	-1.392	-1.429	1.443		3	376	-1.683	-1.747	1.172
FINLAND	1	144	-1.874	-1.969	0.936	IRELAND	1	62	-1.602	-1.520	1.273
	2	72	-1.899	-1.976	0.738		2	31	-1.533	-1.336	1.580
	3	144	-1.047	-1.411	1.570		3	62	-0.844	-0.762	1.438
FRANCE	1	698	-2.028	-1.987	1.028	NETHERLANDS	1	182	-1.678	-1.650	1.032
	2	349	-2.171	-2.195	1.039		2	91	-2.122	-2.206	1.074
	3	698	-1.227	-1.270	1.046		3	182	-0.923	-1.015	1.065
GREECE	1	400	-0.830	-0.906	0.983	SWEDEN	1	368	-1.752	-1.773	1.071
	2	200	-1.156	-1.268	0.885		2	184	-1.828	-1.777	0.996
	3	400	-0.758	-0.917	1.105		3	368	-1.085	-1.098	1.002
ITALY	1	330	-1.529	-1.542	0.939	UNITED	1	722	-1.817	-1.852	1.255
	2	165	-1.809	-1.777	0.920	KINGDOM	2	361	-1.921	-1.850	1.300
	3	330	-0.714	-0.808	0.847		3	722	-1.388	-1.335	1.399
PORTUGAL	1	46	-1.625	-1.816	0.960						
	2	23	-1.972	-1.711	0.965						
	3	46	-0.777	-0.865	1.005						
SPAIN	1	162	-1.436	-1.520	0.965						
	2	81	-1.705	-1.634	1.116						
	3	162	-0.755	-0.842	1.029						

DIFF_GAAP is LARGE (SMALL) if local GAAP differs substantially from IFRS on more than (less than) 12 [i.e., the sample median] of 21 key accounting items as in Bae et al. (2008). Period 1 refers to 2003 and 2004 when IFRS adoption was not mandatory. Period 2 refers to 2005 when IFRS adoption became mandatory. Period 3 covers years after the first year of mandatory IFRS adoption (2006 and 2007).

Table 3: Descriptive Statistics of Selected Variables

Variable	N	Mean	Std	P5	P25	Median	P75	P95
SYNCH(1)	9,520	-2.845	2.136	-7.281	-3.943	-2.407	-1.311	0.137
SYNCH(2)	9,520	-1.587	1.211	-3.457	-2.325	-1.576	-0.797	0.302
SYNCH(3)	9,520	-2.156	1.156	-4.163	-2.854	-2.082	-1.324	0.417
$\text{Beta}_{\text{(MARKET)}}$	9,520	0.467	0.630	-0.489	0.072	0.436	0.840	1.509
$\text{Beta}_{\text{(INDUSTRY)}}$	9,520	0.222	0.476	-0.470	-0.043	0.174	0.465	1.042
Total Volatility	9,520	0.019	0.006	0.012	0.013	0.018	0.024	0.028
NREV	9,520	11.74	24.31	0.00	0.00	1.00	12.00	60.00
INSTIT(%)	9,520	6.53	14.21	0.00	0.00	0.00	5.99	35.59
MCAP	9,520	1254	4793	7	30	112	525	5455
MTB	9,520	2.49	2.72	0.46	1.07	1.73	2.85	6.93
LEVERAGE(%)	9,520	21.81	72.06	0.00	1.19	18.14	39.85	72.76
AGE (in years)	9,520	17.09	9.01	8.00	10.00	14.00	21.00	36.00
HERF	9,520	0.364	0.235	0.067	0.184	0.311	0.496	0.872
NIND	9,520	18.80	21.01	2.00	5.00	11.00	23.00	78.00
GDPG(%)	9,520	2.492	1.268	0.088	1.711	2.600	3.263	4.725
ACCESS	9,520	4.986	0.801	3.740	4.510	4.760	5.560	6.340

Note: All variables are defined in Appendix 1

Table 4: Correlations

Variable	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]	[9]	[10]	[11]	[12]	[13]	[14]	[15]	[16]
[1] SYNCH(2)	1.000	-0.138	0.290	0.300	0.306	0.250	0.298	0.057	0.444	-0.008	0.066	0.102	-0.047	-0.047	0.211	-0.173
		0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.43	0.00	0.00	0.00	0.00	0.00	0.00
[2] ADOPT	-0.151	1.000	-0.360	-0.033	-0.022	-0.457	0.054	-0.056	-0.008	0.019	0.005	0.000	-0.006	-0.011	-0.171	0.000
	0.00		0.00	0.00	0.03	0.00	0.00	0.00	0.43	0.07	0.62	1.00	0.56	0.26	0.00	1.00
[3] POST_ADOPT	0.317	-0.360	1.000	0.108	0.024	0.596	0.025	0.194	0.190	0.020	0.038	0.014	-0.006	-0.061	0.267	-0.021
	0.00	0.00		0.00	0.02	0.00	0.02	0.00	0.00	0.05	0.00	0.17	0.54	0.00	0.00	0.04
[4] $Beta_{(MARKET)}$	0.366	-0.031	0.125	1.000	-0.494	0.060	0.128	0.046	0.124	0.067	0.033	-0.014	-0.002	-0.108	0.054	-0.071
	0.00	0.00	0.00		0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.17	0.82	0.00	0.00	0.00
[5] $Beta_{(INDUSTRY)}$	0.339	-0.016	0.032	-0.427	1.000	0.027	0.098	0.005	0.128	-0.003	0.007	0.026	-0.041	0.121	0.100	-0.096
	0.00	0.11	0.00	0.00		0.01	0.00	0.66	0.00	0.77	0.47	0.01	0.00	0.00	0.00	0.00
[6] Total Volatility	0.276	-0.456	0.589	0.064	0.032	1.000	-0.050	0.081	0.007	-0.028	0.009	-0.110	-0.210	-0.118	0.334	-0.171
	0.00	0.00	0.00	0.00	0.00		0.00	0.00	0.49	0.01	0.38	0.00	0.00	0.00	0.00	0.00
[7] NREV	0.350	0.013	0.087	0.187	0.112	-0.066	1.000	0.123	0.606	0.069	0.067	0.159	0.152	-0.054	-0.051	0.111
	0.00	0.20	0.00	0.00	0.00	0.00		0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
[8] INSTIT(%)	0.103	-0.061	0.254	0.087	-0.002	0.083	0.343	1.000	0.125	0.070	0.004	0.119	0.100	0.012	0.086	0.002
	0.00	0.00	0.00	0.00	0.87	0.00	0.00		0.00	0.00	0.72	0.00	0.00	0.26	0.00	0.82
[9] MCAP	0.458	-0.008	0.191	0.133	0.132	-0.003	0.641	0.197	1.000	0.030	0.146	0.303	0.145	-0.167	-0.014	0.017
	0.00	0.46	0.00	0.00	0.00	0.74	0.00	0.00		0.00	0.00	0.00	0.00	0.00	0.18	0.10
[10] MTB	0.067	0.040	0.065	0.135	-0.027	-0.028	0.289	0.168	0.182	1.000	0.038	-0.064	0.017	0.076	0.035	0.027
	0.00	0.00	0.00	0.00	0.01	0.01	0.00	0.00	0.00		0.00	0.00	0.10	0.00	0.00	0.01
[11] LEVERAGE	0.142	0.001	0.091	0.047	0.042	0.017	0.222	0.025	0.322	0.021	1.000	0.018	0.008	-0.073	-0.011	0.009
	0.00	0.92	0.00	0.00	0.00	0.10	0.00	0.01	0.00	0.04		0.08	0.41	0.00	0.28	0.39
[12] AGE	0.073	0.000	0.020	-0.038	0.008	-0.073	0.079	0.098	0.284	-0.081	0.136	1.000	0.128	-0.100	-0.035	0.012
	0.00	1.00	0.06	0.00	0.44	0.00	0.00	0.00	0.00	0.00	0.00		0.00	0.00	0.00	0.26
[13] HERF	-0.041	-0.005	0.003	0.010	-0.040	-0.204	0.145	0.140	0.127	0.025	0.031	0.060	1.000	0.162	-0.287	0.346
	0.00	0.61	0.77	0.34	0.00	0.00	0.00	0.00	0.00	0.02	0.00	0.00		0.00	0.00	0.00
[14] NIND	-0.050	-0.007	-0.065	-0.109	0.123	-0.116	-0.077	-0.086	-0.132	0.001	-0.159	-0.111	0.197	1.000	-0.128	0.240
	0.00	0.48	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.95	0.00	0.00	0.00		0.00	0.00
[15] GDPG	0.230	-0.166	0.273	0.068	0.102	0.314	-0.030	0.073	-0.005	0.055	-0.056	-0.014	-0.256	-0.131	1.000	-0.407
	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.62	0.00	0.00	0.18	0.00	0.00		0.00
[16] ACCESS	-0.196	0.000	-0.029	-0.082	-0.103	-0.144	0.099	0.090	0.001	0.057	0.011	0.040	0.346	0.210	-0.390	1.000
	0.00	1.00	0.01	0.00	0.00	0.00	0.00	0.00	0.90	0.00	0.30	0.00	0.00	0.00	0.00	

Note: All variables are defined in Appendix 1. Pearson and Spearman correlations are above and below the diagonal, respectively. Two-tailed *p*-values are reported below the correlation coefficients.

Table 5: The Effect of Mandatory IFRS Adoption on Stock Return Synchronicity

Dependent Variable:	SYNCH(1)									SYNCH(2)			SYNCH(3)		
	Market Model									**Industry-Augmented Market Model**			**Fama-French 3-Factor Model**		
	[1]			[2]			[3]			[4]			[5]		
	Parameter Estimate	*t*-stat		Parameter Estimate	*t*-stat		Parameter Estimate	*t*-stat		Parameter Estimate	*t*-stat		Parameter Estimate	*t*-stat	
Constant	-2.030	-9.21	‡	-1.952	-6.44	‡	-1.964	-6.31	‡	-2.123	-10.10	‡	-1.846	-6.81	‡
ADOPT	-0.116	-5.18	‡	-0.108	-4.80	‡	-0.084	-2.81	†	-0.065	-2.29	†	-0.096	-3.19	‡
POST_ADOPT	0.190	6.74	‡	0.230	7.93	‡	0.132	3.76	‡	0.144	4.75	‡	0.181	5.16	‡
$Beta_{(MARKET)}$	0.437	20.56	‡	0.425	20.08	‡	0.426	20.16	‡	0.924	27.92	‡	0.417	19.74	‡
$Beta_{(INDUSTRY)}$										1.209	23.56	‡			
$Beta_{(HML)}$													0.138	8.02	‡
$Beta_{(SMB)}$													-0.098	-4.68	‡
log(Total Volatility)	0.490	10.79	‡	0.487	10.70	‡	0.477	10.46	‡	0.501	12.82	‡	0.437	9.70	‡
log(NREV)				0.087	5.30	‡	0.070	4.13	‡	0.027	1.75	*	0.068	4.06	‡
ADOPT x log(NREV)							-0.008	-0.59		-0.011	-0.92		-0.006	-0.41	
POST_ADOPT x log(NREV)							0.081	5.62	‡	0.045	3.88	‡	0.078	5.43	‡
INSTIT				-0.329	-3.70	‡	-0.344	-3.83	‡	-0.272	-3.67	‡	-0.316	-3.54	‡
ADOPT x INSTIT							-0.272	-1.28		0.063	0.39		-0.273	-1.37	
POST_ADOPT x INSTIT							-0.206	-1.35		0.011	0.09		-0.226	-1.48	
MCAP	0.245	17.04	‡	0.202	16.98	‡	0.198	16.38	‡	0.163	15.04	‡	0.192	16.12	‡
MTB	-0.017	-3.50	‡	-0.016	-3.47	‡	-0.018	-3.71	‡	-0.018	-4.43	‡	-0.016	-3.34	‡
LEVERAGE	0.004	0.44		0.005	0.45		0.006	0.59		0.002	0.20		0.002	0.19	
LOG(AGE)	-0.026	-0.67		-0.006	-0.17		0.001	0.03		0.036	1.14		0.000	0.01	
HERF	-0.006	-1.05		-0.005	-0.79		-0.005	-0.88		-0.004	-0.77		-0.005	-0.80	
Log(NIND)										0.029	1.82	*			
GDPG	0.099	8.87	‡	0.099	8.99	‡	0.098	8.91	‡	0.066	7.29	‡	0.090	8.28	‡
ACCESS	-0.149	-7.71	‡	-0.164	-8.59	‡	-0.167	-8.77	‡	-0.083	-4.95	‡	-0.170	-8.99	‡
N	9,520			9,520			9,520			9,520			9,520		
Adjusted-R^2	0.358			0.365			0.367			0.522			0.374		

Note: Regressions contain industry-fixed effects (1 digit) and standard errors are clustered at the firm-level (Petersen 2009). All variables are defined in Appendix 1. Significance levels ‡ (1%), † (5%) and * (10%).

Table 6: Regression Results of Earnings Informativeness and Stock Return Synchronicity

Dependent Variable:	CAVOL						CAVAR					
	[1]			[2]			[3]			[4]		
	Parameter Estimate	*t*-stat		Parameter Estimate	*t*-stat		Parameter Estimate	*t*-stat		Parameter Estimate	*t*-stat	
Constant	-0.725	-0.78		0.530	0.71		0.527	0.31		1.646	1.06	
SYNCH(2)	-0.334	-3.61	‡				-0.337	-2.32	†			
DR_SYNCH(2)				-1.461	-3.23	‡				-2.094	-4.53	‡
MCAP	0.047	0.86		0.050	0.86		0.231	2.46	†	0.265	2.90	‡
MTB	0.003	0.13		0.002	0.10		-0.022	-0.66		-0.027	-0.81	
D_NEGEPS	-0.210	-0.89		-0.225	-0.97		-0.357	-1.66	*	-0.397	-1.86	*
LEVERAGE	0.030	0.58		0.030	0.58		0.047	0.29		0.046	0.28	
$Beta_{(MARKET)}$	0.591	5.21	‡	0.638	5.21	‡	0.310	1.49		0.501	2.47	†
$Beta_{(INDUSTRY)}$	0.701	4.30	‡	0.746	4.41	‡	0.424	1.56		0.653	2.52	†
Log(NREV)	0.354	6.55	‡	0.359	6.62	‡	0.494	4.94	‡	0.502	5.02	‡
INSTIT	0.335	0.54		0.345	0.56		0.155	0.24		0.092	0.14	
Industry fixed effects	YES			YES			YES			YES		
Country fixed effects	YES			YES			YES			YES		
Firm-clustered s.e.	YES			YES			YES			YES		
N	8,250			8,250			7,940			7,940		
Adjusted-R^2	0.032			0.032			0.055			0.056		

CAVOL is the cumulative abnormal volume for each firm i around the annual earnings announcement date and is measured as the sum of abnormal volume defined as $(V_{id} - \overline{V_i})/\sigma_i$, where V_{id} is equal to the number of shares of firm i traded during day d (d = -1, 0, +1) relative to earnings announcement day (d=0), divided by shares outstanding of firm i during day d; and $\overline{V_i}$ and σ_i are the mean and standard deviation, respectively, of daily trading volume for firm i in the period d-345 to d-20 and d+20 to d+345. CAVAR is the cumulative abnormal volatility measure for each firm around the annual earnings announcement date and is measured as the sum of abnormal volatility defined as u_{id}/σ_i^2, where u_{id} is daily market model-adjusted return computed as $R_{id} - (\alpha_i + \beta_i R_{md})$, R_{id} is the raw return of firm i on day d, R_{md} is equally-weighted return of market for day d, α_i and β_i are firm i's market model parameter estimates, and σ_i^2 is the variance of firm i's market model adjusted returns, each of which is calculated during the period d-345 to d-20 and d+20 to d+345. DR_SYNCH(2) is the scaled decile rank based upon the ranked values of SYNCH(2) and it ranges between zero and one, where higher values represent more synchronicity. D_NEGEPS is equal to one for a firm reporting a loss, and 0 otherwise. All other variables are defined in Appendix 1. Regressions contain industry-fixed effects (1 digit) and standard errors are clustered at the firm-level (Petersen 2009). Significance levels ‡ (1%), † (5%) and * (10%).

Table 7: Regression Results of Stock Return Synchronicity and Mandatory IFRS Adoption for Subsamples Classified by Local GAAP-IFRS Difference

Dependent Variable: SYNCH(2)

	DIFF_GAAP = Large [1]			DIFF_GAAP = Small [2]			Chow (1960) test	
	Parameter Estimate	t-stat		Parameter Estimate	t-stat		F-stat	p-value
Constant	-1.556	-6.44	‡	-2.539	-6.29	‡		
ADOPT	-0.112	-2.84	‡	-0.039	-0.94		5.15	<0.01
POST_ADOPT	0.153	4.45	‡	0.100	1.81	*	11.40	<0.01
$Beta_{(MARKET)}$	1.029	25.41	‡	0.820	16.76	‡		
$Beta_{(INDUSTRY)}$	1.256	18.99	‡	1.126	14.92	‡		
log(Total Volatility)	0.523	11.34	‡	0.460	6.72	‡		
log(NREV)	0.047	3.07	‡	0.005	0.18		2.09	0.12
ADOPT x log(NREV)	-0.005	-0.28		-0.018	-0.99		1.68	0.19
POST_ADOPT x log(NREV)	0.032	2.93	‡	0.048	2.52	‡	7.74	<0.01
INSTIT	-0.276	-2.40	†	-0.119	-1.11		2.80	0.06
ADOPT x INSTIT	-0.410	-1.48		0.299	1.49		2.57	0.08
POST_ADOPT x INSTIT	0.090	0.51		0.047	0.24		0.23	0.79
MCAP	0.146	11.73	‡	0.181	9.95	‡		
MTB	-0.019	-3.44	‡	-0.013	-2.31	†		
LEVERAGE	0.010	0.33		-0.001	-0.12			
Log(AGE)	-0.036	-1.00		0.100	2.14	†		
HERF	0.001	0.10		-0.005	-0.66			
log(NIND)	0.049	2.93	‡	0.019	0.78			
GDPG	0.046	3.70	‡	0.076	4.72	‡		
ACCESS	-0.090	-4.53	‡	-0.118	-1.92	‡		
Industry fixed effects	YES			YES				
Firm-clustered s.e.	YES			YES				
N	4,820			4,700				
Adjusted-R^2	0.574			0.482				

Regressions contain industry-fixed effects (1 digit) and standard errors are clustered at the firm-level (Petersen 2009). All variables are defined in Appendix 1. Significance levels ‡ (1%), † (5%) and * (10%).

Table 8: Regression Results of Stock Return Synchronicity and Mandatory IFRS Adoption for Subsamples Classified by the EPS Restatement Difference Reported At First-time IFRS Adoption

Dependent Variable: SYNCH(2)

	SMALL [1]			MEDIUM] [2]			LARGE [3]		
	Parameter Estimate	t-stat		Parameter Estimate	t-stat		Parameter Estimate	t-stat	
Constant	-1.857	-4.35	‡	-1.973	-5.79	‡	-1.125	-3.01	‡
ADOPT	0.040	0.68		-0.076	-1.35		-0.108	-2.04	†
$Beta_{(MARKET)}$	0.673	18.11	‡	0.809	22.87	‡	0.706	20.87	‡
$Beta_{(INDUSTRY)}$	0.915	17.48	‡	0.975	19.50	‡	1.039	21.73	‡
log(Total Volatility)	0.458	4.92	‡	0.453	5.90	‡	0.586	7.26	‡
Other Controls	YES			YES			YES		
Industry fixed effects	YES			YES			YES		
Firm-clustered s.e.	YES			YES			YES		
N	1,515			1,521			1,515		
Adjusted-R^2	0.462			0.493			0.345		

Note: Regressions use 3 years (2003, 2004 and 2005) of data to estimate the following model for 3 restatement portfolios: $SYNCH_{i,t} = \alpha_0 + \beta_1 ADOPT + \beta_2 Beta_{(MARKET)} + \beta_3 Beta_{(INDUSTRY)} + \beta_4 \log(Total\ Volatility) + \beta_5 Log(NREV) + \beta_6 ADOPT*Log(NREV) + \beta_7 INSTIT + \beta_8 ADOPT_i*INSTIT_i + \beta_9 MCAP + \beta_{10} MTB + \beta_{11} LEVERAGE + \beta_{12} AGE + \beta_{13} HERF + \beta_{14} Log(NIND) + \beta_{15} GDPG + \beta_{16} ACCESS + \gamma_j \sum IND_j + \varepsilon_{i,t}$. For reasons of brevity, we do not report coefficients on control variables other than market and industry characteristics. Firms in the SMALL, MEDIUM, and LARGE restatements portfolio reflect the bottom third, middle third, and top third, respectively, of the restatement measured as the absolute value of 2004 local GAAP earnings per share (EPS) minus the reconciled 2004 IFRS EPS, scaled by local EPS (i.e., $|(EPS_{LOCAL04} - EPS_{IFRS04})/ EPS_{LOCAL04}|$. Regressions contain industry-fixed effects (1 digit) and standard errors are clustered at the firm-level (Petersen 2009). All variables are defined in Appendix 1. Significance levels ‡ (1%), † (5%) and * (10%).

Appendix 1: Variable definitions

Variable	Definition

Dependent Variable

SYNCH(1)

A measure of synchronicity of firm-level stock returns defined as $\log\left(\frac{R^2}{1-R^2}\right)$, where R^2 is the coefficient of determination obtained from the firm-year estimation of the model:

$$RET_{i,w} = \alpha + \beta_1 MARET_{i,w} + \beta_2 MARET_{i,w-1} + \varepsilon_{i,w} \qquad (1)$$

where *MARET* is the value-weighted market return and I, w are firm and week subscripts

SYNCH(2)

A measure of synchronicity of firm-level stock returns defined as $\log\left(\frac{R^2}{1-R^2}\right)$, where R^2 is the coefficient of determination obtained from the firm-year estimation of the industry-augmented market model:

$$RET_{i,w} = \alpha + \beta_1 MARET_{i,w} + \beta_2 MARET_{i,w-1} + \varepsilon_{i,w} \qquad (2)$$

where *MARET* is the value-weighted market return, *INDRET* is the industry value-weighted return using all firms with the same two-digit code, and subscripts i, and w represent a firm and a week.

SYNCH(3)

A measure of synchronicity of firm-level stock returns defined as $\log\left(\frac{R^2}{1-R^2}\right)$, where R^2 is the coefficient of determination obtained from the firm-year estimation of the Fama-French 3-factor model:

$$RET_{i,w} = \alpha + \beta_1 MARET_{i,w} + \beta_1 SMB_{i\ w} + \beta_1 HML_{i,w} + \varepsilon_{i,w} \qquad (3)$$

where *MARET* is the value-weighted market return, *SMB* (small minus big) and *HML* (high minus low) are the weekly returns on zero-investment factor-mimicking portfolios designed to capture size and book-to-market effects, respectively, and subscripts i, and w represent a firm and a week.

Test and Control Variables

ADOPT	Dummy variable equal to 1 if a firm observation is from 2005 fiscal-year (mandatory IFRS adoption year) and 0 otherwise.
POST_ADOPT	Dummy variable equal to 1 if a firm observation is from 2006 or later fiscal-year (Post-mandatory IFRS adoption year) and 0 otherwise.
$Beta_{(MARKET)}$	Firm-year specific market beta estimated (weekly returns). [Datastream]
$Beta_{(INDUSTRY)}$	Firm-year specific industry beta (2-digit SIC) estimated (weekly returns). [Datastream]
$Beta_{(HML)}$	Beta coefficient associated with book-to-market factor in Fama-French (1996) three-factor model. [Bekaert et al. 2009]
$Beta_{(SMB)}$	Beta coefficient associated with size factor in Fama-French (1996) 3 factor model. [Bekaert et al. 2009]
Total Volatility	Standard deviation of weekly market return. [Datastream]
Log(NREV)	Natural log of number of analyst revisions of one-year ahead forecast of annual earnings for the calendar year. [IBES detailed files]
INSTIT	Proportion (expressed as a percentage) of institutional holdings in a firm where institutions are defined as 'pension funds; mutual funds; and insurance companies.' [Amadeus]
MCAP	Market value of equity of the firm, measured at fiscal year-end. [Datastream]
MTB	Market-to-Book ratio [Datastream]
LEVERAGE	Long-term debt divided by (total debt plus shareholders' equity) [Worldscope]
AGE	Number of years since the firm's initial listing date [Datastream]
HERF	A revenue-based Herfindahl index of industry-level concentration. [Worldscope]
Log(NIND)	Natural log of average number of firms used to calculate the weekly industry-level return index. [Worldscope]
GDPG	Inflation-adjusted country growth in gross domestic product [World Bank 2008]
ACCESS	The ease with which firms can raise capital through public equity markets. Scores are expressed as a number varying between 0 (very hard) and 7 (very easy). Values are for 2004 and vary between a minimum of 3.74 (Greece) and a maximum of 6.34 (France). [World Bank 2006]